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Assessing the Contribution of Public Capital to Private Production

Evidence from the German Manufacturing Sector

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Abstract

Using time-series cross-section data from the manufacturing sector of the 11 West German 'Bundesländer' (Federal States) from 1970 to 1996, I examine the impact of public capital on private production. My econometric analysis explicitly takes into account four of the most frequent specification issues in the context of time-series cross-section data analysis: serial correlation, groupwise heteroscedasticity, cross-sectional correlation and nonstationarity of data. For all approaches and tested specifications, I find that public capital is a significant input for production in the manufacturing sector. Moreover, I find that differences in public capital endowment can explain long-term differences in productivity across the Bundesländer. One tentative conclusion that can be drawn from this finding is that differences in public capital endowment might also explain a part of the still-existing productivity gap between manufacturing in East and West Germany. However, I emphasise that the existence of positive effects of public capital on private production is a necessary, but not a sufficient condition for concluding that public investments should be boosted in the future.

1. Introduction

My study is motivated by the controversy that has developed recently about the contribution of public capital—e.g. highways, mass transits, water and sewer systems, etc.—to private production. This controversy has been stimulated by the large elasticity of output with respect to public capital found in the pioneering work of Aschauer (1989b; 1989a). Aschauer's findings suggest that part of the productivity slowdown observed in the 1970s and 80s in the United States and in other OECD countries may be due to an underinvestment in public capital. This has become known in the literature as the 'public capital hypothesis'.

A number of follow-up studies have been spurred by this controversy, some of which have supported the public capital hypothesis (Berndt & Hansson, 1992; Fernald, 1999; Morrison & Schwartz, 1996; Munnell, 1990; Munnell, 1992; Otto & Voss, 1994; Ram & Ramsey, 1989) while others have not (Baltagi & Pinnoi, 1995; Garcia-Milà, McGuire & Porter, 1996; Erber, 1995; Evans & Karras, 1994; Holtz-Eakin, 1994; Hulten & Schwab, 1991; Tatom, 1991; Tatom, 1993).[1] The usual approach taken in these studies is to regress some measure of output e.g. gross domestic product (GDP) or value added on an array of factor inputs and a measure of public capital.[2]

The purpose of this paper is to examine the significance of the 'public capital hypothesis' for Germany. One major finding that emerges from my empirical investigation is that public capital appears to be a significant determinant for private production in the manufacturing sector. Thus, my empirical results are in line with other studies for Germany e.g. Seitz (1993), Licht & Seitz (1994), Seitz (1994) or Schlag (1997). However, I stress that my study (i) uses a different methodology which is not based on the cost but on the less restrictive production function approach, (ii) focuses on the manufacturing sector at the regional level of the Bundesländer and (iii) incorporates several important econometric issues in the statistical analysis which have been neglected in previous studies.

Thus, my study addresses some important methodological concerns raised regarding previous studies. For example, as pointed out by Aaron (1990), Jorgenson (1991) and Tatom (1991; 1993) most of the time series employed for the examination of the relationship between public capital and private output are likely to be nonstationary and thus they advise estimating the model in first differences if the variables are not cointegrated. Following this advice, for instance, Tatom (1991) or Garcia-Milà *et al.* (1996) find the elasticity of output with respect to public capital to be insignificant for the US. This highlights the importance of an appropriate modelling of stochastic or deterministic trends in variables. In my empirical analysis this matter is examined more closely.

Another important motivation of my study is the intention to shed some light onto the nature of the positive correlation between public capital and private output. Thus, I analyse the underlying structure of the data that gives rise to this correlation. The question is whether it results from the variation between cross-sections (Bundesländer) or from the variation over time, i.e. from the 'within' variation. Moreover, I investigate whether this correlation is manifested in the short-run or in the long-run trends in the data.

The remainder of this paper is organised as follows. Section 2 outlines the specification used in the empirical analysis. Section 3 presents the results and considers several econometric specification issues. Section 4 summarises and concludes the paper.

2. Specification

This section considers the specification for my econometric approach to assessing the contribution of public capital to private production.

Suppose that production of value-added output Q_{it} in the manufacturing sector in Bundesland $i = 1, \ldots, B$ at time $t = 1, \ldots, T$ depends on inputs of private capital K_{it} and labour L_{it}. We assume that output Q_{it} also depends on the Hicks-neutral level of technology $A_i()$, which is a function of time t and the level of the non-rival public input G_{it}. Suppose $A_i()$ takes the functional form $A_i = A_{i0} G_{it}^{\beta_g} \exp(\lambda t)$, where A_{0i} is the initial level of technology at time $t = 0$ in Bundesland i and λ is the exogenous rate of technology growth. The exogenous technology growth rate λ is therefore restricted to be the same for all Bundesländer whereas the initial level of technology A_{0i} can vary across the Bundesländer.

Now, specifying a Cobb-Douglas functional form I get the estimating equation in logarithms as

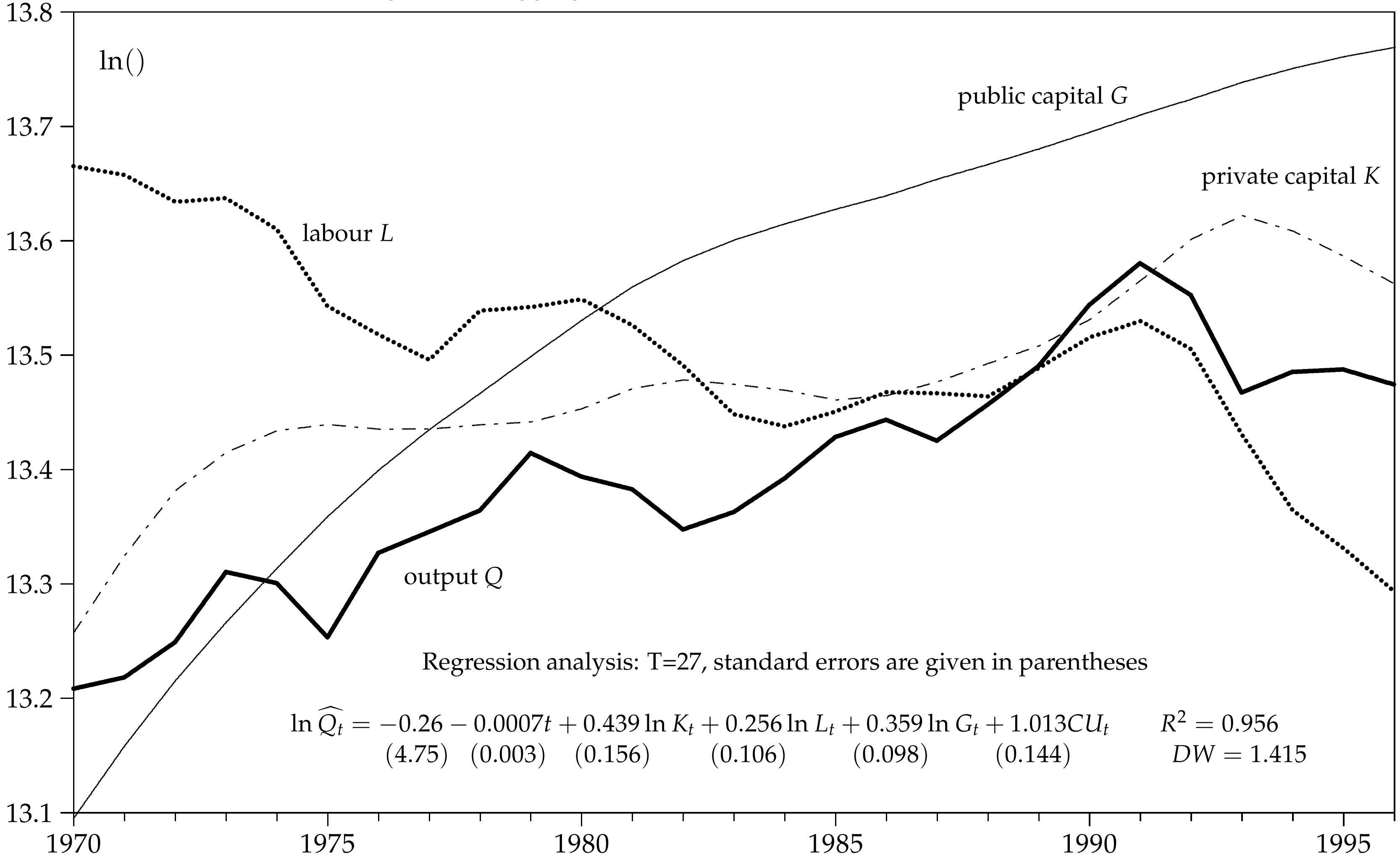
$$\begin{aligned} \ln Q_{it} &= \ln A_{i0} + \lambda t + \beta_g \ln G_{it} + \beta_k \ln K_{it} + \beta_l \ln L_{it} + \beta_{cu} CU + \varepsilon_{it}, \\ i &= 1, ..., B, \qquad t = 1, ..., T, \end{aligned} \tag{1}$$

where ε_{it} denotes an error term which reflects contemporaneous exogenous shocks to logarithmic output $\ln Q_{it}$. We also include a measure for capacity utilisation of private capital in (1), denoted CU.[3] Furthermore, I assume that ε_{it} is an i.i.d. random variable with variance σ_ε^2. Note that in (1) the estimate $\widehat{\beta}_j$, $j \in \{g, k, l\}$, gives the elasticity of output with respect to factor j.

Even if the Cobb-Douglas functional form is restrictive because the elasticities of substitution of input factors are restricted to equal one, it is a first order approximation to any arbitrary production function in the neighbourhood where the factor input vector $X = (G, K, L)$ is $(1, 1, 1)$.[4] It is worth stressing that (1) does not put any restriction on the technology with respect to returns to scale.

It should be mentioned that instead of a production function it would have been possible to specify a dual cost function approach with public capital entering as a quasi-fixed unpaid factor of production. However, at this fairly high level of aggregation the behavioural assumption of the cost function approach that costs are endogenous and determined by choosing cost minimizing quantities of factor inputs given a certain exogenous quantity of output seems to be unrealistic (Berndt, 1991, p. 457). Furthermore, factor prices are quite often not directly observed but have to be calculated using some (restrictive) assumptions which are likely to introduce further sources of measurement error in the data. The production function approach, on the other hand, requires neither a behavioural (minimizing or maximising) assumption nor data on factor prices (Chambers, 1988).

Figure 1: Manufacturing sectors' aggregate series Q, L, K and G for West German Bundesländer from 1970-1996



(For data description and sources, see Appendix)

3. Data, econometric issues and results

3.1. Data

The data used in the analysis cover the manufacturing sector of the 11 West German Bundesländer ($B = 11$) from 1970 to 1996 ($T = 27$). A comprehensive description of the data is given in the Appendix.

Figure 1 graphs the aggregate series of Q, L, K and G as well as CU over the period 1970-1996. Growth of the aggregate public capital stock was particularly high during the period from 1970 to 1981. After 1983 the growth rate of public capital declined slightly compared to the previous period, but was still positive and relatively constant.

On the other hand the aggregate private capital stock in manufacturing grew at a relatively high rate from 1970 to 1975, but in the period 1976-1985 the growth rate of the capital stock slowed. Note that changes in aggregate private capital appear to follow changes in output with a lag of about two to three years. For example the decrease in output during the years 1979 to 1982 seems to have had an effect on the formation of private capital after 1982. Hence, at least at the aggregate level, there is some evidence that private capital formation is likely to follow the development in output and not vice versa. Similarly, from 1993 to 1996 I observe a decline in the stock of private manufacturing capital whereas a sharp decline in output occurred already from 1991 to 1993.

The aggregate labour series shows a clear downward trend over the total period. This can be ascribed to the structural change in the German economy where the share of the manufacturing sector's employment in the total economy is declining.

Note also that the series of output and labour show rather high annual fluctuations due to the business cycles of the economy, whereas the series for capital, and in particular for public capital, are smoother. One reason for this is that planning and decisions in public investments are oriented toward the long term, sometimes with a horizon of five to 15 years. Therefore, annual fluctuations in output, i.e. fluctuations due to business cycles, do not appear to have an impact on the short-term formation of public capital. However, in the long run, business cycles are likely to influence the formation of public capital due to the effects of the business cycles on tax revenues.

In addition, Figure 1 also presents the results of a regression analysis where output Q is regressed on inputs L, K, G, capacity utilisation CU and a linear time trend t. The basic specification for the estimation is an autoregressive model of order one (AR(1)), which has been estimated by applying the iterated Prais-Winsten method (Greene, 2000, p. 547).

We find that labour $\ln L_t$, private capital $\ln K_t$, public capital $\ln G_t$ as well as capacity utilisation CU are statistically significant at a five percent level, whereas the linear time trend t is not. The fit of this preliminary regression with a R^2 of about 0.95 is remarkably high. It is worth mentioning that the estimate for labour with a value of 0.26 appears to be too low with regard to the share of wages in value added of the manufacturing sector in my sample, which is about 0.55.

3.2. *Basic model results for inputs K and L*

To begin with the main part of the empirical analysis based on the pooled time-series cross-section data, I first present results for the model where only private inputs are included in (1), i.e. K_{it} and L_{it}. This preliminary step is undertaken in order to be able to evaluate the changes in results due to the inclusion of the public capital input G_{it} in the production function (1). In the second step I therefore present estimation results for the model with all inputs, including G_{it}.

Ordinary Least Squares (OLS) estimation of the Cobb-Douglas production function as specified in (1) with private inputs K_{it} and L_{it} using the pooled time-series cross-section data yields the following estimates:[5]

$$\begin{array}{lllllll} \widehat{\ln Q_{it}} = & \text{Länder-effects}^{\star} & +0.015^{\star}\, t & +0.134\, \ln K_{it} & +0.672^{\star} \ln L_{it} & -0.096^{\star} CU & \\ & F(10|283)=142.9 & (0.001) & (0.078) & (0.082) & (0.187) & \end{array} \quad (2)$$

N: 297 $(G = 11, T = 27)$ $R^2 : 0.9932$ $SE : 0.867$

Diagnostic test

Test for serial correlation: $DW : 0.266^{\star}$, $\rho_{LM} = 209.2^{\star} \sim \chi^2_{df=1}$

Test for groupwise heteroscedasticity: $LM = 146.6^{\star} \sim \chi^2_{df=10}$

Test for cross-sectional correlations: $\lambda_{LM} = 447.4^{\star} \sim \chi^2_{df=55}$

Test for random walk of residual: $R_p : 0.165$

Hausman test: 2.36

Multicollinearity: condition number = 495.4

Notice that in (2) the included dummy variables for the Bundesländer ('Länder' effects) correspond to the term $\ln A_{i0}$ in (1). The displayed F-test indicates that these Bundesländer effects are highly significant. The value of 2.36 of the Hausman test favours a random effects model against the fixed effects model. Furthermore, labour is significant with a value of 0.672. However, the estimate of private capital is not significant at a five percent level. Note that the fit of the regression with R^2 equal to 0.9932 is remarkably high.

A frequent observation in the empirical analysis of time-series data is the presence of autocorrelation. Also, it is very likely that heteroscedasticity will be observed as the Bundesländer in my sample differ in size. Furthermore, macroeconomic factors affecting one region will also affect other regions, thus the errors across the Bundesländer are likely to be correlated.

3.3. *Specification tests*

Autocorrelation

In order to explore these econometric specification issues, several diagnostic checks are shown in (2).[6] First, to test for the presence of autocorrelation, the value of the Durbin-Watson (DW) statistic, which is 0.266, has been calculated from the residuals of the OLS estimation according to Bhargava, Franzini & Narendranathan (1982) as

$$DW = \frac{\sum_{i=1}^{G}\sum_{t=2}^{T}(\widetilde{u}_{it} - \widetilde{u}_{i,t-1})^2}{\sum_{i=1}^{G}\sum_{t=1}^{T}\widetilde{u}_{it}^2},$$

where $\widetilde{u}_{it}$ are the residuals from the fixed effects model (2). The Durbin-Watson statistic can be used to test the null hypothesis that the serial correlation is $\rho = 0$ against the alternative that $|\rho| < 1$. The *exact* critical value for the DW statistic is 1.810 and has been found by using the Imhof (1961) routine.[7] Thus, the null that the errors of the OLS estimation are serially independent is rejected.

This finding is also confirmed by the value of the Lagrange-Multiplier test statistic $\rho_{LM} = 209.2$.[8] This statistic is distributed χ^2 with 1 degree of freedom ($\chi^2_{crit,0.05,df=1} = 3.84$), hence I can reject the null hypothesis of serial independence at a five percent level by this test.

Groupwise heteroscedasticity

Second, in order to test for groupwise heteroscedasticity the following Lagrange multiplier (LM) test has been calculated as (Greene, 2000, p. 596)

$$LM = T/2 \sum_{1}^{B} \left[\frac{s_i^2}{s^2} - 1 \right]^2 ,$$

where s^2 is the pooled OLS residual variance and s_i^2 is the estimated unit-specific residual variance from groupwise regressions. The LM statistic has a limiting χ^2 distribution with $B - 1$ degrees of freedom. The reported value of 146.6 from the LM statistic leads to a rejection of the null hypothesis of no groupwise heteroscedasticity ($\chi^2_{crit,0.05,df=10} = 18.3$).

Cross-sectional correlations

Third, in order to test for cross-sectional correlations the residuals obtained from (2) are used to compute the following Lagrange multiplier statistic (Greene, 2000, p. 452)

$$\lambda_{LM} = T \sum_{i} \sum_{j<i} r_{ij}^2,$$

where r_{ij}^2 is the squared ijth correlation coefficient of residuals between Bundesland i and j. The large-sample distribution of this statistic is chi-square with $B(B-1)/2$ degrees of freedom. Hence, this statistic with a value of 447.4 is

Table 1: Cross-sectional correlation and variance/covariance[1] matrix for the 11 Bundesländer based on residuals from equation (3)

	BaW	Bay	Ber	Bre	Ham	Hes	Nie	NRW	RhP	Saa	SHo
BaW	**1.56**	1.06	2.30	-0.06	-0.57	1.49	1.01	1.06	0.78	2.21	1.15
Bay	0.78	**1.18**	2.94	-1.23	-2.34	1.25	0.83	0.90	0.69	2.14	-0.14
Ber	0.49	0.72	**14.21**	-12.30	-14.90	2.16	0.47	1.79	0.65	6.51	-0.27
Bre	-0.01	-0.29	-0.83	**15.52**	14.75	0.64	1.90	0.14	1.01	-3.02	0.10
Ham	-0.09	-0.42	-0.78	0.74	**25.90**	-1.30	1.87	-0.67	-0.24	-5.71	5.80
Hes	0.86	0.83	0.41	0.12	-0.18	**1.92**	1.32	1.24	1.04	2.43	0.07
Nie	0.58	0.55	0.09	0.35	0.26	0.68	**1.93**	1.02	0.80	1.40	0.12
NRW	0.86	0.84	0.48	0.04	-0.13	0.90	0.74	**0.97**	0.73	1.81	0.11
RhP	0.64	0.66	0.18	0.26	-0.05	0.77	0.59	0.76	**0.94**	1.16	-0.32
Saa	0.80	0.89	0.78	-0.35	-0.51	0.80	0.46	0.83	0.54	**4.85**	-0.03
SHo	0.36	-0.05	-0.03	0.01	0.44	0.02	0.03	0.04	-0.13	-0.01	**6.64**

BaW=Baden-Würtemberg, Bay=Bayern, Ber=Berlin, Bre=Bremen, Ham=Hamburg, Hes=Hessen, Nie=Niedersachsen, NRW=Nordrhein-Westfalen, RhP=Rheinland-Pfalz, Saa=Saarland, Sho=Schleswig-Holstein

[1]Variances / covariances [10^{-3}], correlations are given below, covariances above and variances in bold on the diagonal

highly significant, indicating the presence of substantial cross-sectional correlations between the Bundesländer ($\chi^2_{crit,0.05,df=55} = 73.3$).

Table 1 shows the correlations r_{ij} and variances / covariances of residuals between the Bundesländer. The variances of the residuals of the Bundesländer are given in bold print on the diagonal of the matrix. Covariances are given in the upper half of Table 1. The ratio of the largest variance with 25.9 ('Hamburg') to the smallest with 0.94 ('Rheinland-Pfalz') is about 27, which confirms the high degree of groupwise heteroscedasticity in the data. Similarly, some of the correlations shown in the lower half of Table 1 are remarkably high, for instance between 'Baden-Württemberg' and 'Hessen' with a value of about 0.86.

Stationarity

Fourth, as a first glance at Figure (1) revealed that the (aggregate) series exhibit some (random or deterministic) trends, the R_p statistic[9] according to Bhargava *et al.* (1982) for testing the null that the residuals from (2) follow a random walk, i.e. $\rho = 0$ against $|\rho| > 0$, is also presented. Small values of R_p favour the null hypothesis. The exact critical value for this statistic again can be found by using the Imhof routine. In my case the critical value for R_p at a five percent level is 0.336[10], therefore the null hypothesis of a unit root is not rejected.[11] Since the error is nonstationary, the variables appear not to be cointegrated.

Multicollinearity

Finally, the paper by Ai & Cassou (1997) points out that the findings of some studies for the US using fixed effects models in the analysis of productivity effects of public capital based on state level data, e.g. Holtz-Eakin (1994) or Evans & Karras (1994), should be interpreted with some caution because of the high correlation between the public capital stocks and the fixed effects. This multicollinearity problem arises because there is not enough variation in the public capital series to disentangle the effect of public capital from the state-specific effect, i.e. the public capital series do not have enough 'within' variation. Thus, to get some indication whether multicollinearity matters for my estimations I also report the condition number[12] which has a value of 495.4. Judge, Griffiths, Hill, Lee & Lütkepohl (1985, p. 902) suggest that values exceeding 20 reveal potential multicollinearity problems. Thus, the occurrence of poor or imprecise estimates can be a result of the high degree of multicollinearity in the data.

3.4. Basic model results for all inputs K, L and G

Estimating (1) for all inputs, i.e. K_{it}, L_{it} and G_{it} I obtain the following results:

$$\widehat{\ln Q_{it}} = \underset{F(10|282)=154.9}{\text{Länder-effects}^*} + \underset{(0.002)}{0.002}\,t - \underset{(0.064)}{0.106}\,\ln K_{it} + \underset{(0.065)}{0.753^*}\,\ln L_{it} + \underset{(0.058)}{0.779^*}\,\ln G_{it} + \underset{(0.147)}{0.045^*}\,CU \tag{3}$$

N: 297 $(G = 11, T = 27)$ $R^2 : 0.9958$ $SE : 0.805$

Diagnostic tests

Test for serial correlation: $DW : 0.389^\star$, $\rho_{LM} = 191.7^\star \sim \chi^2(df = 1)$

Test for groupwise heteroscedasticity: $LM = 147.2^* \sim \chi^2(df = 10)$

Test for cross-sectional correlations: $\lambda_{LM} = 537.4^* \sim \chi^2(df = 55)$

Test for random walk of residuals: $R_p : 0.230$

Hausman test: 18.91^*

Multicollinearity: condition number = 594.5

Again, I find that the coefficient of labour input is significant, whereas the coefficient of private capital is not. In contrast to this, the estimate of the coefficient of public capital input is highly significant. Here, the value of 18.91 of the Hausman test favours the fixed effects model against a random effects model. Also, from the increase in the Hausman test statistic from 2.36 in (2) to 18.91 in (3) I infer that public capital appears to be correlated with the Bundesländer effects. Hence, the random effects model should not be applied.

The displayed diagnostic tests reveal that all the specification issues for estimation such as serial correlation, groupwise heteroscedasticity and cross-sectional correlation are present as before. Again, the null hypothesis of a random walk of the residuals is not rejected at a five percent, since the R_p statistic does not exceed the critical value of 0.336.

3.5. *Estimation strategy*

Our further estimation strategy is therefore as follows. From the reported R_p statistics in (2) and (3) respectively it is generally difficult to judge whether a trend stationary or difference stationary model is more appropriate. In the former case the estimation can be carried out in levels, whereas for the latter case the estimation should be based on variables in first differences. Therefore, I will present estimation results both for the specification in levels and for the specification in first differences. This also allows us to check the robustness of the results obtained.

Additionally, instead of calculating robust PCSEs[13] due to groupwise heteroscedasticity and cross-sectional correlation, another estimation strategy is to apply Feasible Generalised Least Squares (FGLS) in order to properly take into account serial correlation, groupwise heteroscedasticity and/or cross-sectional correlation. Feasible Generalised Least Squares (FGLS) estimation in the context of time-series cross-section models is also known as the 'Kmenta' or 'Parks' method (Kmenta & Oberhofer, 1974; Kmenta, 1986; Parks, 1967). Beck & Katz (1995) have argued that one should be aware of the fact that although FGLS might be more efficient when cross-sectional correlations or groupwise heteroscedasticity are very significant, the standard errors obtained by the FGLS estimation do not correctly reflect the sampling variability of parameter estimates, because in samples of small size the cross-sectional correlations or variances obtained in the first step of FGLS are likely to be very poor estimates of the underlying 'true' variances. Thus, as Beck & Katz (1995) have shown by Monte-Carlo simulations, standard errors from FGLS estimation in small samples have a tendency to be too small, they are 'overconfident'. Beck & Katz recommend applying OLS estimation with consistent and robust panel corrected standard errors (PCSE) instead of FGLS if the ratio of number of time periods to the number of cross-sections is smaller than three. This is the case for my sample, since the ratio of T to B is 2.45. Thus, there is a risk that standard errors obtained from FGLS are 'overconfident'. Therefore, I present results both for FGLS estimation as well as for OLS with PCSEs.

For the AR(1) models a consistent estimate of the autocorrelation parameter ρ was obtained from residuals of equation (2) and (3) respectively as $\widehat{\rho} = 1 - \mathrm{DW}/2$. Using this estimate, the first step AR(1) correction has been carried out by employing the Cochrane-Orcutt transformation (Greene, 2000, p. 546). As such, the first observation in each group is lost.[14] In the second step, I use two estimation variants. The first variant—which is, due to the AR(1) correction in the first step, also an FGLS estimation—is based on OLS estimation in the second step with robust panel corrected standard errors (PCSEs) of the transformed variables. The second variant is based on FGLS estimation in both steps ('Kmenta' method)— in the first step an AR(1) correction is performed and, in the second step, the FGLS estimation is performed taking groupwise heteroscedasticity and cross-sectional correlation into account. Also, for the model in first differences I apply both estimation methods, i.e. (i) OLS with PCSEs and (ii) FGLS ('Kmenta' method).

3.6. Empirical results

Table 2 summarises the results of the estimations. The upper half (I) contains the results for inputs K and L, and in the lower half (II), the results for inputs K, L and G. Columns 1 and 2 present the results for the AR(1) models, whereas columns 3 and 4 display the results for variables in first differences. Note that only the AR(1) models include the Bundesländer dummy variables (fixed effects), since the dummy variables are removed when taking first-differences. Similarly, only the AR(1) models include a time trend t, because the time trend becomes a constant when taking first-differences.

The usual F-test for OLS relies on homoscedasticity. If this is not an appropriate assumption one can use a Wald test instead.[15] Both the F-tests and the Wald tests show that the Bundesländer effects are highly significant. Also, the null hypothesis of constant returns to scale (CRS) is rejected in almost all specifications, but not in (II) for all inputs K, L and G.

By contrast with the low values for the DW statistics reported for the previous estimations (2) and (3), both the AR(1) and the model with variables in first differences generate DW statistics above 1.810, indicating that autocorrelation and also stationarity of residuals are not problematic for the estimations. This is further confirmed by the Lagrange-Multiplier statistic ρ_{LM}, which does not reject the null hypothesis of serial independence for most specifications at a five percent level.

The parameter estimates of private capital, labour input, public capital and capacity utilisation are significant in all specifications. The estimate for labour input with values between 0.248 and 0.498 appear to be somewhat too low considering again that the average (wage) share of labour in output in my sample is about 0.55. Notice also the decrease of the condition number from the AR(1) to the specification in first differences. Hence, for the specification in first differences multicollinearity is no longer problematic for the estimations.

Table 2: Production function estimates (G=11, T=27)

(I) Dependent variable $\ln Q_{it}$, factors of production K_{it}, L_{it}

	AR(1), $\rho = 0.869$				**first differences**			
	FGLS (PCSE)		FGLS (Kmenta) (het., corr.)		OLS (PCSE)		FGLS (Kmenta) (het., corr.)	
const	Länder-dum.*		Länder-dum.*		0.009*	(0.003)	0.013*	(0.002)
t	0.001	(0.004)	0.008*	(0.002)	—	—	—	—
$\ln K_{it}$	0.368*	(0.139)	0.218*	(0.052)	0.431*	(0.141)	0.321*	(0.061)
$\ln L_{it}$	0.248*	(0.094)	0.452*	(0.048)	0.270*	(0.132)	0.375*	(0.058)
CU	0.841*	(0.135)	0.748*	(0.055)	0.812*	(0.143)	0.819*	(0.060)
R^2	0.940		—		0.400		—	
N	286		286		286		286	
Diagnostic tests								
	F tests:		Wald tests χ^2:		F tests:		Wald tests χ^2:	
fixed effects	9.28*		290.1*		—		—	
CRS K, L	13.64*		7.81*		8.74*		4.11*	
DW	1.898		—		1.924		—	
LM test: ρ_{LM}	0.252		—		0.17		—	
cond.-number	199.6		—		2.58		—	

(II) Dependent variable $\ln Q_{it}$, factors of production K_{it}, L_{it}, G_{it}

	AR(1), $\rho = 0.783$				**first differences**			
	FGLS (PCSE)		FGLS (Kmenta) (het., corr.)		OLS (PCSE)		FGLS (Kmenta) (het., corr.)	
const	Länder-dum.*		Länder-dum.*		-0.001	(0.006)	0.006*	(0.003)
t	-0.005	(0.004)	0.004*	(0.002)	—	—	—	—
$\ln K_{it}$	0.298*	(0.133)	0.105*	(0.052)	0.269*	(0.099)	0.257*	(0.066)
$\ln L_{it}$	0.256*	(0.119)	0.498*	(0.045)	0.301*	(0.089)	0.421*	(0.058)
$\ln G_{it}$	0.651*	(0.217)	0.547*	(0.100)	0.537*	(0.134)	0.385*	(0.092)
CU	0.825*	(0.094)	0.717*	(0.057)	0.825*	(0.094)	0.825*	(0.094)
R^2	0.972		—		0.432		—	
N	286		286		286		286	
Diagnostic tests								
	F tests:		Wald tests χ^2:		F tests:		Wald tests χ^2:	
fixed effects	14.40*		354.9*		—		—	
CRS K, L	19.65*		10.86*		17.1*		8.46*	
CRS K, L, G	1.55		1.36		0.57		0.02	
DW	1.827		—		2.045		—	
LM test: ρ_{LM}	1.090		—		0.462		—	
cond.-number	642.6		—		3.74		—	

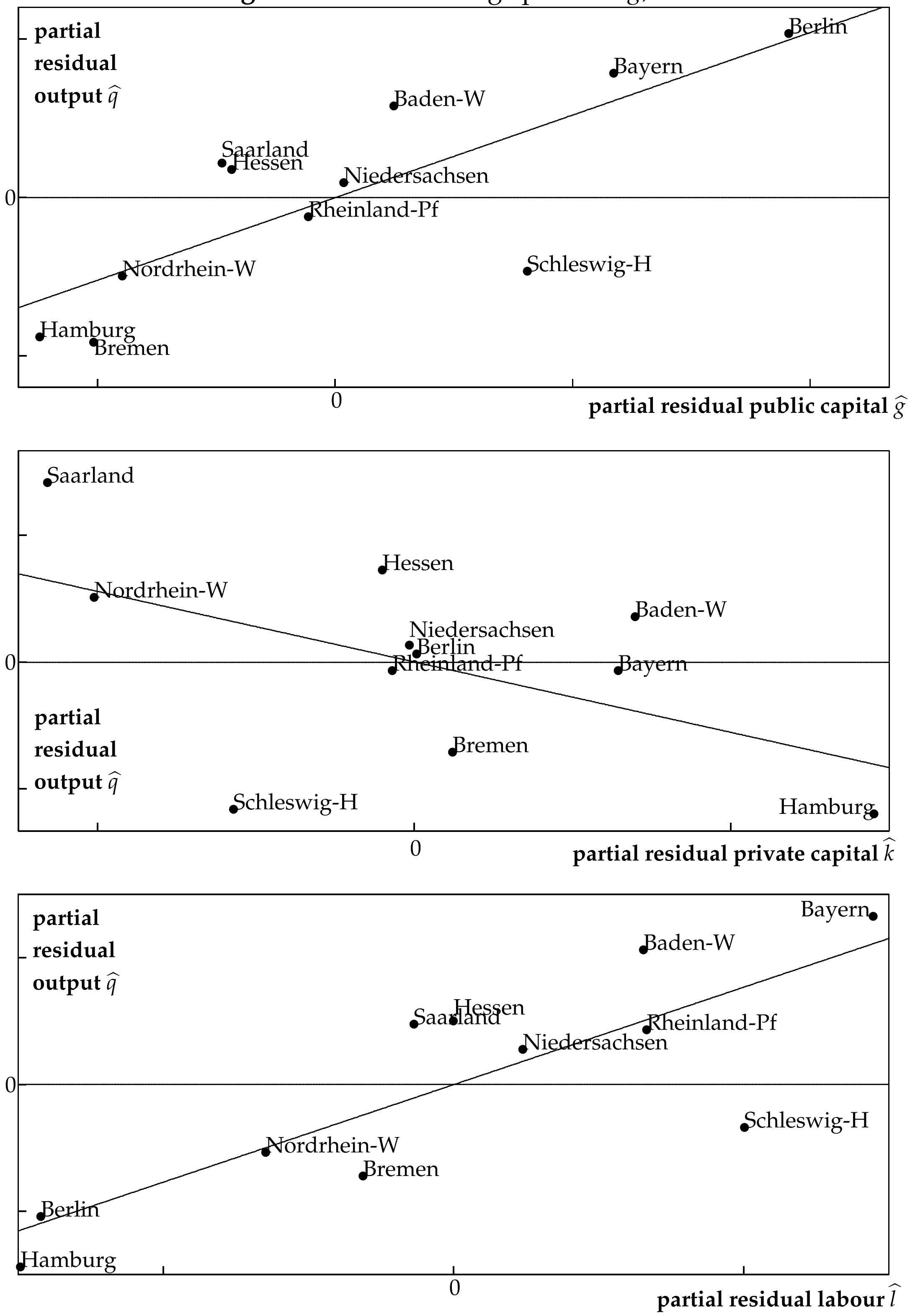
* denotes statistical significance at a 5 % level, standard errors are given in parentheses

The results in (II) show that in contrast to private capital, the parameter for public capital appears to be significant in all specifications with values ranging between 0.38 and 0.65. Since the ratio of output Q to public capital G varies between 1.12 in year 1970 and 0.69 in year 1996, these estimated elasticities imply a marginal productivity of G between 43 and 73 percent in 1970 and between 26 and 45 percent in 1996.[16] The differences in R^2 between (I) and (II) are about 0.03. Hence, in my model the public capital input can explain about three percent of the differences in observed output across the Bundesländer.

Stability of parameter estimates and poolability

Finally, I provide several tests on the stability of parameters both over (i) cross-sections ('testing for poolability of the data') and (ii) over time ('testing for occurrence of structural breaks in the data'). In order to test (i) I perform a Chow test (Baltagi, 1995, chap. 4.1) on the null hypothesis that the parameters (including the intercept) across the Bundesländer are equal, i.e. $H_0 : \beta_i = \beta, i = 1 \ldots B$. To accomplish this, based on the model in first differences from Table 1, an observed $F-$value of 0.929 is obtained which is distributed as $F(50, 251)$ under the null. This does not reject poolability across the Bundesländer. Similarly, for testing (ii), $H_0 : \beta_t = \beta, t = 1 \ldots T$, an observed F-value of 1.841 is obtained. Note that a structural break, i.e. a change of the parameter vector over time, can only be significant if at least one of the parameter vectors β_t differ from β. Based on the central $F(75, 182)$ distribution, the null that the parameters across time can be pooled is rejected at the one percent level.[17] However, if we are willing to trade some bias for a reduction in variance, some weaker criteria can be used (Baltagi, 1995, p. 54). The null hypothesis is then that the restricted model is better than the unrestricted model in terms of the trade-off between bias and variance. As a criterion for this test I use the noncentrality parameter λ (Baltagi, 1995, p. 55). From the observed λ value of 0.38, the null hypothesis is neither rejected by the first and second 'weak' MSE criterion (Wallace, 1972) nor by the 'strong' MSE criterion (Toro-Vizcarrondo & Wallace, 1968). Thus the pooling of the time-series cross-section data is supported.

Figure 2: Partial leverage plots for $\widehat{g}$, $\widehat{k}$ and $\widehat{l}$



Correlation structure

As the final step of my empirical analysis, in order to shed some light on the underlying structure of the positive correlation between public capital and output, I consider a very simple regression where the growth rate of output in the period 1970-1996, denoted by $\widehat{q}$, is regressed on the growth rates of inputs denoted by $\widehat{k}$, $\widehat{l}$, $\widehat{g}$, over the same period.

The first regression with only inputs $\widehat{k}$ and $\widehat{l}$ yields the following result:

$$\begin{array}{rccc} \widehat{q}_i = & 0.584 & -0.423\,\widehat{k}_i & +0.568\,\widehat{l}_i \\ & (0.187) & (0.378) & (0.291) \end{array} \tag{4}$$

N: 11 $\quad R^2 : 0.369 \quad F : 2.34$

The second regression with inputs $\widehat{k}$, $\widehat{l}$ and $\widehat{g}$ gives:

$$\begin{array}{rcccc} \widehat{q}_i = & 0.075 & -0.416\,\widehat{k}_i & +0.768\,\widehat{l}_i & +0.867\,\widehat{g}_i \\ & (0.196) & (0.251) & (0.202) & (0.259) \end{array} \tag{5}$$

N: 11 $\quad R^2 : 0.758 \quad F : 7.29^*$

The two regressions show that long-term changes in public capital and labour (in the 'between' Bundesländer dimension) are associated with long-term changes in manufacturing sector's outputs. The coefficients of labour and public capital are in line with the previous results, although the estimates are somewhat higher. Note that over a long period such as this, capacity utilisation is negligible for realised output.

Figure 2 presents the partial leverage plots for regression (5). Two reference lines are displayed in the plots. One is the horizontal line where the partial residual of $\widehat{q} = 0$, and the other is the fitted regression of the partial residual of $\widehat{q}$ against the partial residual of the respective input.[18] The latter has an intercept of 0 and a slope equal to the parameter estimate associated with the explanatory variable in the model.[19] The partial leverage plots reveal that the results of the regression (5) are not driven by single influential observations. Except for Schleswig-Holstein all observations contribute positively to the partial correlation between $\widehat{q}$ and $\widehat{g}$. Also, it can be seen from Figure 2 that the insignificance of $\widehat{k}$ is not determined by single influential observations. Interestingly, both Hamburg and the Bundesländer Saarland and Nordrhein-Westfalen, which experienced the most intense structural change in the manufacturing sector with strongly declining heavy industries during the last two decades, do not fit into a hypothetical positive partial correlation between private capital and output. This hints that the low significance of private capital could be driven by the structural change in the manufacturing sector which made large parts of the private capital stock obsolete.

We also performed a further regression which is not reported here where the average level of output was regressed on the average levels of the inputs over the period 1970-1996. Thus, the number of observations for this 'between' regression is again 11. It turned out that parameters of all inputs were insignificant. Hence, from this evidence I conclude that differences in levels of public capital or in public capital intensity, defined as the ratio of public capital to labour, do not matter for differences in productivity across the Bundesländer. This is not a surprising finding considering that the level of public capital endowment for each Bundesland also depends on the geographical characteristics of the Bundesland.[20]

3.7. Summary of main findings

From the econometric analysis of this section the following three key findings of this study can be recorded. First, and most important, the stylised finding of this study is that public capital is significant for production in the manufacturing sector. This holds for all tested econometric models and specifications. For variables in levels, this result is mainly driven by the 'within' variation whereas the 'between' variation does not contribute to it. Thus, differences in public capital intensity can not explain differences in observed levels of output, but differences in changes of public capital can explain differences in changes of output. Furthermore, this correlation between changes of public capital and output holds both in the short-run and in the long-run dimension.

Second, differences in public capital growth can explain about three percent of the differences in the manufacturing sector's output growth across Bundesländer over the period 1970 to 1996.

Third and finally, at least for the sample studied here, the factor inputs and output appear not to be cointegrated series. For the model with all inputs, i.e. labour L, private capital K and public capital G, the model in first differences appears to give the most reliable results.

4. Conclusions

The starting point of this paper has been Aschauer's (1989a,1989b) public capital hypothesis, which states that the decline in government's infrastructure spending in the US and other major OECD countries during the 1970s and 80s can explain a major part of the observed decline in productivity growth over the same period.

Several methodological improvements to related studies have been incorporated into the analysis in this paper. We have explicitly taken into account four of the most frequent specification issues in the context of time-series cross-section analysis: serial correlation, groupwise heteroscedasticity, cross-sectional correlation and nonstationarity of the data. Furthermore, I have used a specification in the analysis that has avoided a potential simultaneity problem between output and factor inputs. Finally, I have provided tests on the poolability of data and the stability of parameters over time.

In summary, I find a strong positive and significant correlation between public capital and the manufacturing sector's output at the regional level of the Bundesländer in all of the tested specifications.

One tentative conclusion that can be drawn from this finding is that differences in public capital endowment might also explain a part of the still-existing productivity gap between manufacturing in East and West Germany. Recent studies (Komar, 2000; Seidel & Vesper, 2000) report that the gap in public capital endowment on a per capita basis between East and West German regions is still about 30 percent, while at the same time productivity of firms located in East Germany is only about two-thirds of the productivity of firms located in the West. Thus, at least a part of the productivity differences might be also attributed to differences in public capital endowment.

Given the significance of public capital for private production, one potential economic policy question is whether the process of convergence in public capital endowment between East and West German regions should be accelerated over the next years. At this point, however, I emphasise that the existence of positive effects of public capital on private production is a necessary, but not a sufficient condition for drawing the conclusion that public investments should be boosted in the future. To make this inference, the costs of financing the public capital provision have to be included in the analysis as well. For instance an increase in public investments may only be possible if tax revenues are also increased. This in turn can give rise to distortions bearing additional costs for the economy. Similarly, if higher public investments are financed by higher governmental debt, this may also imply other kinds of additional costs e.g. higher interest rates on capital markets. In this respect, my study has focused only on the necessary condition for increasing the supply of public capital, i.e. the existence of significant and positive effects of public capital on private production. In a more rigorous fashion, the sufficient condition for increasing public investments is that the social net benefit—defined as the sum of social gross benefits (consumer and producer surpluses, positive externalities e.g. spillover effects, etc.) minus the sum of social costs (costs of provision, negative externalities e.g. environmental effects, etc.)—has to be positive.

The obtained estimates of the output elasticity of public capital between 0.38 and 0.65 imply rate of returns between 26 (minimum) and 72 (maximum) percent for my sample. Since these are measures for the return of public capital only for manufacturing, but do not capture the returns for other economic sectors, they appear to be too high to be a plausible estimate of the 'true' returns of public capital for manufacturing. On the other hand these magnitudes are in line with other studies which have been also conducted for the manufacturing sector e.g. Morrison & Schwartz (1996).

A fundamental problem of both my study and related ones is that there is no certainty whether or not other factors that might also positively contribute to the manufacturing sector's output have been omitted from the analysis. If these factors are positively correlated with public capital but excluded in the regression equation, then the expected value of the parameter of public capital will be upward biased. Such a factor could be for instance the stock of knowledge or of the available technology in the manufacturing sector. However, it very difficult to find plausible measures for these intangible stocks, since they are not directly observable.

A promising line for future research is to compare the outcomes of the production, dual cost and profit function approaches as in Vijverberg, Vijverberg & Gamble (1997), who use time-series data for the US and do not find much agreement between the three approaches. The main advantage of this research strategy is that it opens the avenue to study whether the obtained results are robust with respect to the applied (dual) methodology.

Notes

This article is a substantial revision of chapter one of my 2001 dissertation at Humboldt-University Berlin (Stephan, 2001). I am grateful to Charles Blankart, Michael Burda, Almas Heshmati, Astrid Jung, Ulrich Kamecke, Lars-Hendrik Röller, seminar participants at United Nations University (WIDER) in Helsinki and the editor of this journal for helpful comments and suggestions. I also thank Deborah Bowen for proof reading the manuscript. All remaining errors or omissions are solely the author's responsibility.

[1]For comprehensive surveys on this literature, see for instance Gramlich (1994), Sturm, Kuper & de Haan (1996) or Pfähler, Hofmann & Bönte (1997).

[2]Berndt & Hansson (1992), Erber (1995) and Morrison & Schwartz (1996) are based on a dual cost function instead of a production function. Hulten & Schwab (1991) and Fernald (1999) use total factor productivity (TFP) growth as the dependent variable in the analysis.

[3]This measure is only available at the aggregate level, see data description in the Appendix.

[4]In order to capture the second order effects I also estimated flexible functional forms for the production function e.g. translog (Christensen, Jorgenson & Lau, 1971; Christensen, Jorgenson & Lau, 1973) in the empirical analysis. However, it turned out that the estimation of these specifications suffered from a strong multicollinearity problem. This problem arises from extremely high correlations of the single factor inputs with the quadratic and the cross effect terms.

[5]$\star$ denotes statistical significance at a five percent level

[6]In all these cases, OLS estimation still yields consistent parameter estimates. However, estimates of standard errors will be biased and inconsistent.

[7]The Imhof routine is implemented in SHAZAM 8.0.

[8]The Lagrange multiplier statistic is found by regressing $\widetilde{u}_{it}$ on $\widetilde{u}_{i,t-1}$ and the other regressors. The statistic ρ_{LM} is then defined as the R^2 obtained from this auxiliary regression multiplied with the number of observations.

[9]The Rp statistic is calculated as $Rp = e'e/e'F^{*}e$, where e are the OLS residuals from estimating (2) in first differences, $F^{*} = (I_G \otimes F)$, and F is a $(T-1) \times (T-1)$ symmetric matrix with elements of the form $F_{jk} = (T-j)k/T$ if $j \geqq k$ and $F_{jk} = F_{kj}$.

[10]The 10 percent critical value is 0.307, the 1 percent critical value for Rp is 0.398 ($B = 11, T = 27$).

[11]For alternative approaches of testing for unit roots with panel data, see for instance Baltagi & Kao (2000) or Maddala & Kim (1998).

[12]The condition number is defined as the square root of the ratio of the largest to the smallest Eigenvalue of $S(X'X)S$, where S is a diagonal matrix with the kth diagonal element $1/\sqrt{x_k' x_k}$.

[13] Robust panel corrected standard errors (PCSEs) are given in parentheses. PCSEs are estimated by the square root of the diagonal of $(X'X)^{-1}X\,(\Sigma \otimes I_T)\,X(X'X)^{-1}$ where Σ is a $N \times N$ matrix of cross-sectional variances and covariances. A consistent estimate of Σ is given by $E'E/T$, where E denotes $T \times i$ matrix of OLS residuals from equation (2) (Beck & Katz, 1996).

[14]Note that in the fixed-effects model, the Prais-Winston transformation (Greene, 2000, p. 546) is not an appropriate choice for an AR(1) correction, because the 'within' transformation, that is computing deviations from group means, will not remove the heterogeneity if the Prais-Winston transformation is used for the first observation.

[15]The Wald statistic W is defined as $W = (\mathbf{R}\boldsymbol{\beta} - \mathbf{q})' \left(\mathbf{R}(\mathrm{Var}(\widehat{\boldsymbol{\beta}})\mathbf{R}'\right)^{-1} (\mathbf{R}\boldsymbol{\beta} - \mathbf{q})$, where $\mathbf{R}\boldsymbol{\beta} = \mathbf{q}$ imposes a set of restrictions on the parameter vector $\boldsymbol{\beta}$ (Greene, 2000, p. 361).

[16]The rate of return r_G is obtained from the estimated elasticity as $r_G = \beta_G Q/G$.

[17]Note that CU is not included in this poolability test because it does not possess cross-sectional variation.

[18]The partial residual of $\widehat{q}$ is obtained by regressing $\widehat{q}$ on $\widehat{k}$ and $\widehat{l}$. The partial residual of an input is obtained by regressing this input on the other inputs.

[19]The leverage plot also shows the changes in the residuals for the model with and without the explanatory variable. For a given data point in the plot, its residual without the explanatory variable is the vertical distance between the point and the horizontal line; its residual with the explanatory variable is the vertical distance between the point and the fitted line.

[20]Note also that three of the Bundesländer, i.e. Berlin, Bremen and Hamburg, are agglomerated urban Bundesländer which have very different public capital intensities compared to the territorial Bundesländer.

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Appendix

A Data[21]

Output Q

Output is measured as gross value-added at market prices of the manufacturing sector in 1991 constant prices aggregated over industries in year *t*. These data have been obtained from the series *'National accounts for the Bundesländer'* (engl. transl.), issue 30: 'Entstehung des Bruttoinlansprodukts in den Ländern der Bundesrepublik Deutschland 1970 bis 1996', which is provided by the Statistical Office of Baden-Würtemberg. For years 1991-1996 I obtained updated figures (in mid 2000) from the Statistical Office of Baden-Würtemberg. These updated figures had not yet been published.

Public capital (G)

Public capital is measured as the public gross stock of fixed assets at the ground level ('Bruttoanlagevermögen öffentlicher Tiefbau') at the end of year *t*. It is given in 1991 constant prices. It includes capital formation of all levels of government in Germany, i.e. the local governments, the Federal States ('Bundesländer') governments and the Federal Government. The main parts of this stock are roads and highways (about 50 percent), bridges and railways, but also water and sewer systems, dikes and ports, etc. Note that these stocks are measured according to international convention in gross terms because of the very low depreciation rate for these types of fixed assets.

The figures for the public gross stock of fixed assets have been provided by the Statistical Office of Baden-Würtemberg from the study group of the 'National accounts of the Bundesländer' and have not yet been published.

Private capital (K)

Private capital is measured as the net stock of fixed assets in the manufacturing sector at the end of year t in constant prices of 1991. It includes machinery, equipment and construction, and is taken from '*National accounts for the Bundesländer*', issue 29: 'Anlageinvestitionen, Anlagevermögen und Abschreibungen in den Ländern der Bundesrepublik Deutschland 1970 bis 1995'. This statistical report is also provided by the Statistical Office of Baden-Würtemberg from the study group of the 'National accounts of the Bundesländer'. For years 1991-1996 I obtained revised and updated figures from the Statistical Office of Baden-Würtemberg.

Labour (L)

Labour is measured as the number of employees in the manufacturing sector at the regional level of the Bundesländer. These data have been drawn from the series 'Statistical Yearbook for the Federal Republic of Germany' published by the Federal Statistical Office in Wiesbaden. These figures are measured each year after the first quarter on the 1st of April. Thus to estimate the value at the end of year t I have computed $3/4 * L_t + 1/4 * L_{t+1}$ as a weighted average for years t and $t + 1$.

Alternatively to this labour input measure, I have also estimated the production function with the number of working hours (only of blue-collar employees, also given for the 1st of April) as the labour input which I obtained from the same publication mentioned above. The differences in the obtained parameter estimates are rather small, therefore I have refrained from reporting these results.

Wages (L)

Wages cover both blue- and white-collar employees in the manufacturing sector at the regional level of the Bundesländer. The date of reference is the 1st of April for each year. These data have been obtained from the series 'Statistical Yearbook for the Federal Republic of Germany' published by the Federal Statistical Office in Wiesbaden. In the empirical analysis, wages are only used to compute the average share of labour in total income. For my sample, this share is about 55 percent.

Capacity utilisation (CU)

Capacity utilisation of private capital in manufacturing was only available at the aggregate level and has been obtained from the business survey of the IFO institute in Munich. We have also tried to compute regional level CU measures. To accomplish this, I have regressed the aggregate level measure on two regional level proxies of CU, i.e. the manufacturing's usage of electricity and regional material tons transportation, both measures available from the Federal Statistical Office in Wiesbaden. However, I found that the obtained measures for regional CU gave dissatisfactory results in the following estimations, because the coefficient of private capital became negative. Therefore I refrained from using regional proxy measures of CU.